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Statistical modeling of the volume-outcome effect for carotid endarterectomy for 10 years of a statewide database

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Objective: **We aimed to achieve accurate statistical modeling of a putative relationship between carotid endarterectomy (CEA) annual surgeon and hospital volume and in-hospital mortality.**
Design of Study: **We performed a secondary data analysis of 10 years (1994-2003) of the Maryland hospital discharge database. Annual volume was defined as the total number of procedures performed for the time in the dataset divided by the total years in the dataset. Non-linear relationships between death and average volumes were explored with logit-transformed lowess smoothing functions, followed by random effect models and inspection of data likelihood under each combination of spline knots. A marginal model with generalized estimating equations was used to represent population-average response as a function of covariates and to account for clustering in the data. Patient comorbidity was assessed using the Deyo modification of the Charlson Index.**
Setting: **The Maryland hospital discharge database is a 100% sample of all hospitals in the state.**
Subjects: **CEA was identified through ICD-9 and diagnosis codes, using a previously reported algorithm.**
Main Outcome Measure: **Estimated odds ratios predicting in-hospital death, α set at 0.05.**
Results: **During the study period, 22,772 patients with surgeon identifiers underwent CEA in Maryland, resulting in 123 in-hospital deaths (0.54%). The crude odds ratio of death for the entire surgeon dataset was 0.9838, meaning that the odds of death decreased by an average of 0.0162 for each additional annual procedure. Surgeon volume of four to 15 CEAs per year was highly significant: for an increase in annual surgeon volume by one procedure per year, the estimated odds of death decreased by 0.065 when controlling for hospital volume, age, and comorbidity ($P = .351$). Surgeons in other volume categories also demonstrated lower odds of death with increased annual volume, but these odds ratios did not attain statistical significance. Surgeons performing ≤3 CEA per year had an odds ratio of death of 0.802 per additional annual procedure ($P = .351$), whereas those performing >15 CEAs per year had an odds ratio of 0.997 ($P = .485$). Hospitals that saw >130 CEAs per year had an odds ratio of death of 0.945 per additional procedure, or 0.055 decrease in the odds of death ($P = 0.013$), whereas hospitals performing ≤130 CEAs per year had an odds ratio of 0.998 ($P = 0.563$).**
Conclusion: **We have demonstrated a technique for rigorous statistical analysis of volume-outcome data and have found a volume effect for death after CEA in this 10-year Maryland dataset. Higher volume surgeons had lower estimated odds of death, particularly those performing four to 15 CEAs per year. These data suggest that a patient undergoing CEA by a surgeon performing an average of 16 CEAs annually has a statistically equivalent risk of death compared with one undergoing CEA by a surgeon performing any number higher than this, when controlling for hospital volume, patient comorbidity, and patient age. Hospital volume was not seen to be as significant a predictor of postoperative death in this study, with only high volume hospitals (≥130 CEAs per year) showing a statistically significant decrease in the odds ratio of death. As studies on volume-outcome relationships can have important implications for health policy and surgical training, such studies should consider non-linear effects in their modeling of procedural volume. (J Vasc Surg 2008;48:343-50.)**

Carotid endarterectomy (CEA) remains one of the most common vascular surgical procedures in the United States. More than 50 years after its implementation by Eastcott and Robb, CEA remains the "gold standard" treatment for carotid artery disease.[1-3] The predominance of this procedure has been established by numerous studies demonstrating its safety and efficacy. The North American Symptomatic Carotid Endarterectomy Trial (1991)[4] and The Asymptomatic Carotid Atherosclerosis Study (1995)[5] supported the superiority of the CEA over the best medical management of patients with ≥50% symptomatic or ≥60% asymptomatic carotid stenoses, respectively.

Despite the proven efficacy of this procedure, questions remain as to whether all patients have access to the high quality surgeons and hospitals reflected in these and other

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pivotal studies. Studies of other surgically treated diseases support the notion that "practice makes perfect": those surgeons or hospitals performing the greatest number of surgeries yearly provide the best outcomes.[6-17] Numerous studies of CEA have established tiers of surgeons or hospitals performing low-, middle-, and high-numbers of CEAs per year and demonstrated differences in outcomes among these strata.[15,18-24] Using this evidence, practitioners and health policymakers have argued that higher volume surgeons and hospitals render superior CEA results to patients, leading to calls for regionalization of CEA care.[25-27]

To date, investigators of the volume-outcome relationship for CEA surgery have arrived at their thresholds for high-, medium-, and low-volume surgeons and hospitals either empirically or based on quantiles.[18-22,24,28] While many authors have found a volume effect for CEA, their wide-ranging definitions for volume categories makes it difficult to effect definitive policy changes or for patients to know what annual surgical volume does make a difference to their outcome.

We hypothesized that we could establish evidence-guided volume guidelines for hospital and surgeon volume using standard statistical techniques. By allowing the data of 10 years of the Maryland hospital discharge database to drive the analysis, we hypothesized that we could find the best-supported volume categories for CEA. Furthermore, we hypothesized that a model including other covariates such as age, race, gender, and comorbidities would establish the relative contribution of surgeon and hospital volume compared with these other factors. Because of the state-specific nature of our dataset, our hypotheses focused on the effects of recent surgeon and hospital experience, rather than "lifetime" experience.

MATERIALS AND METHODS

The present study was a secondary data analysis of 10 years (1994-2003) of the Maryland Health Services Cost Review Commission (HSCRC) database. Established in 1971 by the Maryland legislature, this organization aims primarily to contain rising medical costs. In the process, the HSCRC collects extensive information on all patients and medical procedures occurring within the state of Maryland. We analyzed the HSCRC dataset over the abovementioned 10-year period, using a previously reported algorithm to identify 22,772 patients undergoing CEA.[29-31] Patients with all three of the following diagnoses were included: 1) International Classification of Diseases, 9th Revision, Clinical Modification (ICD-9-CM) procedure code 38.12 (endarterectomy of the vessels of the head and neck other than intracranial vessels) in the primary coding position but not in any secondary position, and 2) Diagnosis-Related Group (DRG) 5 (extracranial vascular procedure), and 3) Diagnosis code 433.00 to 433.91 (occlusion/stenosis, precerebral artery). The accuracy of this algorithm was confirmed through a previously published chart review, in which the presence of all three of the abovementioned codes was required to achieve 100% correct identification of CEA patients.[29] Patients receiving CEA concurrently with another procedure, such as coronary artery bypass grafting, were not included in this analysis, based on the previous performance of the algorithm.

We analyzed the relationship between in-hospital death and annual surgeon and hospital volume. Annual surgeon volume was defined by the total number of procedures performed by a surgeon for their total time within the dataset divided by the number of years the surgeon was included in the dataset. Thus a surgeon who was included in the dataset for only five years would have an average volume that could be compared with any other surgeon who was included for any amount of time between one and ten years.

Crude odds ratios of death were first determined by logistic regression for annual surgeon volume and annual hospital volume. Heterogeneity by calendar year was explored by performing the analysis within each year. Nonlinear relationships between death and average annual surgeon and hospital volume were explored by examining logit-transformed lowess smoothing functions. A lowess smoothing function facilitates visual inspection of the data as well as quantitative assessment by fitting a weighted linear least squares regression over small localized segments of the data. Thus a scatterplot of data is rendered more visually accessible by the superimposition of a smooth line that can reflect slope changes across the range of the explanatory variable. If the lowess smoothed line has a natural inflection point (ie visually noticeable change in the slope), the investigator can then elect to treat different ranges of the explanatory variable individually in the analysis, by allowing the slope to change over the range of the variable. In other words, an inflection point (known as a "knot" statistically) can be used to demarcate two different slopes in the regression. Since the outcome of our analysis was odds of death, we used a curve smoothed by log odds ("logit").[32] Rough identification of spline knots from the plots was followed by using random effect models and inspecting likelihood of the data under each combination of spline knots. When various combinations of knots appear possible from preliminary visual inspection of the data, the investigator may wish to weigh the explanatory power of including all possible knots versus more parsimonious models including only a subset of knots. Random effects models allow each subject (each surgeon in this case) to have a different intercept such that minimal assumptions are made about correlations between surgeons. Consideration was given to other values around the knots initially identified, ± 5 CEAs per year. This exploration of values around the knots identified by visual inspection was done to ensure that the best value had been chosen in terms of its explanatory power. A functional form for age was similarly determined by inspection of logit transformed lowess smoothed plots.

Analysis proceeded similarly for looking at hospital volumes. Annual average hospital volume was defined analogously as for annual average surgeon volume: number of procedures performed in the hospital in the complete

10-year dataset, divided by the number of years the hospital was included in the dataset.

The effect of patient co-morbidity was determined through use of unweighted Charlson comorbidities.[33] The Charlson Index was developed in 1987 as a way to communicate the risk profile of a patient by pooling weighted values for selected conditions. The weights and chosen comorbidites were developed by Charlson et al using an inpatient cohort of 604 patients admitted to the medical service of New York Hospital. We chose to disregard the weights in the present study, instead using the number of Charlson comorbidities as identified by ICD-9-CM codes in the dataset, per Deyo.[34] The list of comorbidities is as follows:

Myocardial infarct	Hemiplegia
Congestive heart failure	Moderate or severe renal disease
Peripheral vascular disease	Diabetes with end organ damage
Cerebrovascular disease	Any tumor
Dementia	Leukemia
Chronic pulmonary disease	Lymphoma
Connective tissue disease	Moderate or severe liver disease
Ulcer disease	Metastatic solid tumor
Mild liver disease	AIDS
Diabetes	

Adapted from.[33]

Patients with ≥3 Charlson comorbidities were used as a reference in comparison with patients with one or two comorbidities for association with risk of death.

A marginal model with generalized estimating equations (GEE) was used to represent the population-average response as a function of covariates and to account for clustering in the data. Clustering was accounted for only at the surgeon level, not hospital.

All measures of statistical significance were based on $\alpha = 0.05$. Stata SE, version 10, from Stata Corporation (College Station, Tex) was used for data analysis. The final model employed general estimating equations using exchangeable correlation.

RESULTS

From 1994-2003, CEA was performed on 22,772 patients in Maryland, in 47 hospitals by 442 surgeons. This sample consisted of 54.7% men and 45.3% women, ranging in age from 33-99 years (mean, 70.6 years). The vast majority of the patients were white (21,229 or 91.4%); 1,682 were black (7.2%). There were 123 in-hospital deaths (0.54%) over this 10-year period among those with surgeon identifiers.

Examination of the histogram of average annual surgeon volume demonstrates a strong predominance of low-volume surgeons (Fig 1). This distribution of CEAs among surgeons follows a roughly inverse power relationship, with 214 of the surgeons (48.42%) performing an average of only one CEA per year.

The crude odds ratio of death for the entire surgeon dataset was 0.9838, meaning that the odds of death decreased by an average of 0.0162 for each additional annual procedure.

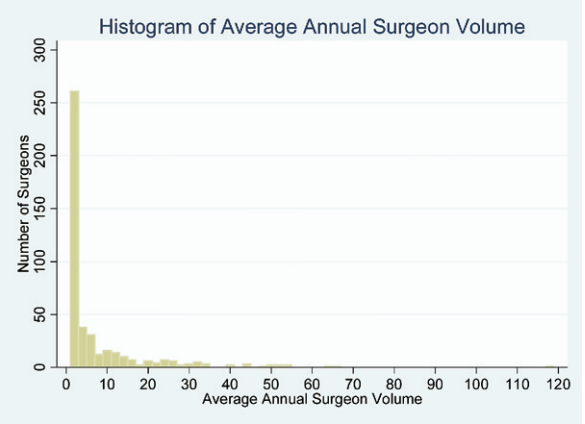

Fig 1. Histogram demonstrating the distribution of annual CEA volume among the surgeons in the dataset.

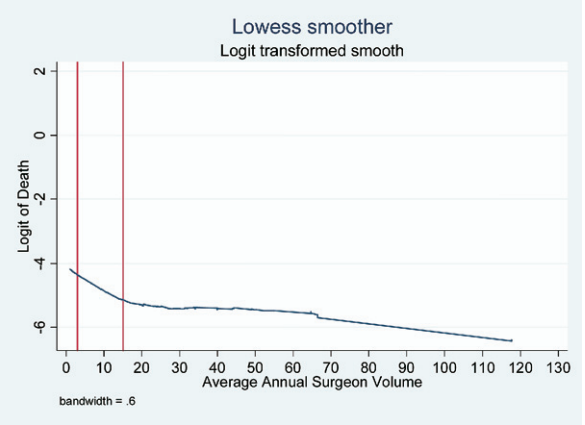

Fig 2. Logit transformed lowess smoothed curve of annual CEA surgeon volume.

The possibility of heterogeneity of volume effect by calendar year was considered. However, since the number of observations (deaths following CEA) per calendar year was small, it was not possible to detect heterogeneity if it did exist. For the interpretation of this study, we assumed homogenous effect of surgeon volume on death over calendar years.

Examination of the logit (log-odds) transformed lowess smoothed functions suggested three knots for surgeon volume: around five CEAs per year, 20 CEAs per year, and 35 CEAs per year. (Fig 2). Using random effects models we examined log likelihoods of different models around these knots (i.e. 3, 4, 5, 15-25, 30-40, etc.) and found the highest log likelihood to be rendered by knots at 3 CEAs per year and 15 CEAs per year. Despite the suggestion of a knot around 35 CEAs per year, the *P* value for this knot was not significant. The lowess smoothed plot for patient age clearly suggested an inflection at 60 years, a value that was

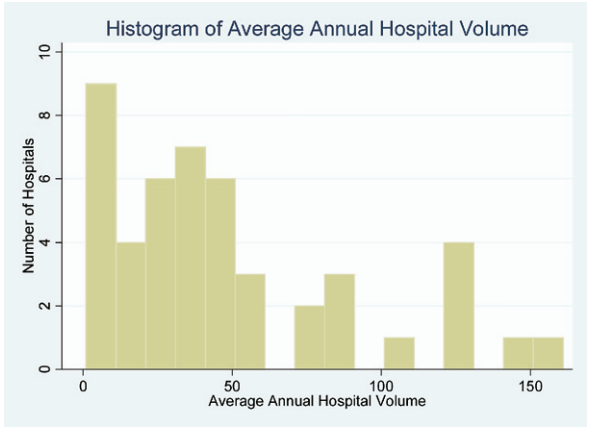

Fig 3. Histogram demonstrating the distribution of annual CEA volume among the hospitals in the dataset.

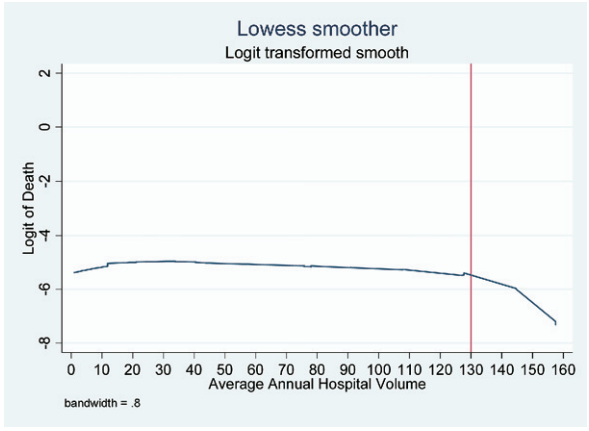

Fig 4. Logit transformed lowess smoothed curve of annual CEA hospital volume.

subsequently borne out by examination of ages grouped around 60 years. Patients over age 60 comprised 85.96% of the sample.

In analysis of hospital volume, a high proportion of low-volume hospitals was again noted in histogram analysis (Fig 3). Examination of the logit transformed smooth curve suggested two splines at annual hospital volume of 25 CEAs per year and 130 CEAs per year (Fig 4). Examination of values around 25 CEAs per year did not support inclusion of this knot in the model, but the knot at 130 CEAs per year was retained due to statistical significance.

This final model was used to predict the odds of in-hospital death following CEA:

$$\begin{aligned}\log it(death_{ij}) &= \beta_0 + \beta_1(surgeonvolume) \\ &+ \beta_2(surgeonvolume)^{3+} + \beta_3(surgeonvolume)^{15+} \\ &+ \beta_4(hosptialvolume) + \beta_5(hosptialvolume)^{130+} \\ &+ \beta_6(age) + \beta_7(age)^{60} + \beta_8(1comorbidity) \\ &+ \beta_9(2comorbidities)\end{aligned}$$

Logistic regression rendered the following point estimates, 95% confidence estimates and P values (Table I). As can be seen in the bolded values, surgeon volume of four to 15 CEAs per year was highly significant with respect to odds of death. For an increase in annual surgeon volume by one procedure per year, the odds of death decreased by 0.065 when controlling for hospital volume, age and number of Charlson comorbidities. In other words, a patient whose surgeon performs six CEAs per year would be expected to have 0.935 the odds of death of a patient whose surgeon performs five CEAs per year, averaged across the population. The same inference can be drawn for any one CEA per year difference in annual surgeon volume in this volume range of four to 15 CEAs per year. Hospitals that saw >130 CEAs per year had an odds ratio of death of 0.945 per additional procedure annually ($P = .013$), or 0.055 decrease in the odds of death. The impact of each additional year of age on odds of death after CEA was U-shaped. For each additional year in age for those ≤60 years, the odds ratio per year of age was 0.936, but statistically non-significant ($P = .115$). However, patients >60 years had an odds ratio of death of 1.058 per additional year of age ($P < .0001$). Finally, the number of Charlson comorbidities was highly significant, with each additional comorbidity adding significantly to the odds of death after CEA. Those with a single Charlson comorbidity had an odds ratio of death of 0.187 ($P < .0001$) compared with those with three or more comorbidities, while those with two comorbidities had an odds ratio of 0.362 ($P < .0001$) compared with the reference.

The inclusion of gender in the model did not change any of the inferences and was itself insignificant ($P = .7715$). Female gender had an odds ratio of death of 0.948 compared with men when controlling for the covariates above. Non-white race (black and "other") had a higher odds of death after CEA than whites when controlling for other factors, but this relationship was non-significant.

Mortality rates for each volume group are shown in Table II, and cross-tabulation of mortality rates for subgroups of surgeon volume categories within hospital volume categories are shown in Table III. These tables include cumulative numbers of surgeons and CEAs over the 10-year study.

DISCUSSION

In this analysis of predictors of death for 22,772 CEAs with surgeon identifiers performed over 10 years in Maryland, the following covariates were found to be significant: 1) Each additional annual procedure for surgeons performing four to 15 CEAs per year (odds ratio 0.935, $P = .013$); 2) Each additional annual procedure for hospitals performing >130 CEAs per year (odds ratio 0.945, $P = .013$); 3) Each additional year of age >60 (odds ratio 1.058, $P < .0001$); 4) One Charlson comorbidity (odds ratio of 0.187 compared with three or more Charlson comorbidities, $P < .0001$); 5) Two Charlson comorbidities (odds ratio of 0.362 compared with three or more Charlson comorbidities, $P < .0001$).

Table I. Odds ratios of death following CEA for predictors included in the model

Predictor	*Odds ratio of death*	*95% Confidence limits*		P *value*
Surgeon Volume ≤3*	0.802	0.505	1.275	.351
Surgeon Volume 4-15*	0.935	0.887	0.986	**.013**
Surgeon Volume >15*	0.997	0.987	1.006	.485
Hospital Volume ≤130*	0.998	0.993	1.004	.563
Hospital Volume >130*	0.945	0.904	0.998	**.013**
Age ≤60^	0.936	0.863	1.016	.115
Age >60^	1.058	1.030	1.088	**<.0001**
1 Charlson Comorbidity`	0.187	0.109	0.322	**<.0001**
2 Charlson Comorbidities`	0.362	0.232	0.566	**<.0001**
≥3 Charlson Comorbidities`	Reference			

*Odds ratio of death per additional procedure per year within volume category.
^Odds ratio of death per additional year of age within age category.
`Odds ratio of death for number of Charlson comorbidities compared with reference of ≥3.

Table II. Number of surgeons and hospitals in different volume categories, with corresponding numbers of CEAs and deaths

Volume category	*Number in category (percent)*	*Number of CEAs in category (percent)*	*Number of deaths (mortality rate)*
Surgeon volume 1	214 (48.4)	248 (1.1)	5 (2.0)
Surgeon volume ≤3	60 (13.6)	447 (2.0)	7 (1.6)
Surgeon volume 4-15	109 (24.7)	5,729 (25.2)	42 (0.7)
Surgeon volume >15	59 (13.4)	16,348 (71.8)	69 (0.4)
Hospital Volume ≤130	45 (95.7)	19,749 (86.7)	119 (0.6)
Hospital Volume >130	2 (4.3)	3,023 (13.3)	4 (0.1)

Table III. Numbers of surgeons within hospital categories, with corresponding numbers of CEAs and deaths

Hospital volume category	*Surgeon volume category*	*Number in category (percent)*	*Number of CEAs in category (percent)*	*Number of deaths (mortality rate)*
Hospital volume ≤130		45 (95.7)	19,749 (86.7)	119 (0.6)
	Surgeon volume 1	198 (49.0)	230 (1.2)	5 (2.2)
	Surgeon volume ≤3	55 (13.6)	417 (2.1)	7 (1.7)
	Surgeon volume 4-15	98 (24.3)	5,099 (25.8)	41 (0.8)
	Surgeon volume >15	53 (13.1)	14,003 (70.9)	66 (0.5)
Hospital volume >130		2 (4.3)	3023 (13.3)	4 (0.1)
	Surgeon volume 1	16 (42.1)	18 (0.6)	0 (0.0)
	Surgeon volume ≤3	5 (13.2)	30 (1.0)	0 (0.0)
	Surgeon volume 4-15	11 (29.0)	630 (20.8)	1 (0.2)
	Surgeon volume >15	6 (15.8)	2,345 (77.6)	3 (0.1)

Odds ratios measure effect size, and describe the odds of an event occurring in one group to the odds of it occurring in another group. Generally speaking, odds represent the probability (p) of an event occurring to the probability of it not occurring: odds = $(p/1-p)$. Thus an odds ratio of 1 indicates that an outcome (in-hospital death after CEA in this study) is equally likely between two groups. Logistic regression, employed here, provides odds ratios per unit change in the predictor. Thus an estimated odds ratio of 0.935 for CEA annual volume four to 15 means that a patient undergoing CEA by a surgeon performing five CEAs annually would have 0.935 the odds of in-hospital death compared with his odds of death should he have had a surgeon performing four CEAs annually, other covariates being equal. This difference may appear small at first glance, but magnifies as one compares slightly different annual surgeon volumes. To compare the odds of death for undergoing CEA by a surgeon performing nine CEAs per year to one performing four CEAs per year, one exponentiates the odds ratio: $0.935_5 = 0.715$. Therefore, when accounting for age, hospital annual volume, and comorbidities, the same patient undergoing CEA by a surgeon performing nine CEAs annually would be expected to have less than ¾ the risk of death compared with the risk should she undergo the surgery by a surgeon performing four CEAs annually.

The decrease in the odds of death per additional annual surgeon procedure differs depending on the volume range.

Our study suggests that the difference per additional procedure is even more pronounced for annual surgeon volumes less than or equal to three. The odds ratio supported by our data for this volume range is 0.802, although this did not achieve statistical significance ($P = .351$).

A similar effect was seen in analysis of the effect of annual hospital volumes on the odds of death, but only in high volume centers. For the majority of hospitals in the database (those with annual CEA volume ≤130), the volume effect was exceedingly small and statistically non-significant per additional annual procedure (odds ratio 0.998, $P = .563$). However, a statistically significant volume effect was seen in higher-volume centers, those performing >130 CEAs per year. In these higher volume centers, the odds ratio of death per additional procedure was 0.945, $P = .013$. As before, this effect amplifies when considering large differences in annual volume within the range: a patient undergoing surgery at a center performing 150 CEAs per year would have an odds of death of 0.945^{20} or 0.323 compared with undergoing surgery at a center that saw 130 CEAs per year. However, this finding should be interpreted with caution. There were very few centers with >130 annual CEAs included in this dataset and therefore these findings may not be generalizable. Observed differences in the odds of death could be due to center-specific factors, rather than annual CEA volume. In spite of the suggestion of an inflection point around 25 CEAs per year on the hospital volume spline, no knots in this range were found to be significant. Thus the apparent rise in slope for the smoothed log odds of death in the low volume hospital range was not substantiated in likelihood ratio analysis. Although previous studies have found profound volume-outcome effects for hospital volume, they may not have controlled for surgeon volume as we did in this analysis. Similarly, our findings for high-volume surgeons are limited by the smaller number of providers in the upper ranges annual volume. Of note, there were 59 surgeons over the 10-year study (13.4% of our sample) in our category of >15 CEAs per year. Analysis of the effect on odds of death of an additional procedure per year for this group did not yield a statistically significant effect ($P = .485$). It is possible that a larger sample size in this volume category could have rendered this effect significant. It is notable, however, that for both surgeons and hospitals, higher annual volume was always predictive of lower odds of death, even if the inference did not attain statistical significance. In no case did we find higher odds of death with higher annual CEA volume.

We found a U-shaped relationship for odds of death relative to age in this study. Each additional year of age in patients ≤60 years led to a decreased risk of death but did not attain statistical significance (odds ratio 0.936, $P = .115$). However, in patients >60 years, each additional year contributed to the odds of death after CEA (odds ratio 1.058, $P < .0001$). We expect this U-shaped relationship is due to different risk profiles in the young undergoing CEA compared with those of advanced age. Young patients undergoing CEA may have risks that were not captured by our efforts to control for comorbidity in this analysis.

Finally, the presence of comorbidities proved to be highly significant in contributing to risk of in-hospital death following CEA. We used ≥3 Charlson comorbidities as the reference group, and compared the odds of death for patients with one or two comorbidities. After controlling for annual surgeon volume, hospital volume and age (≤60 years versus >60 years), patients with only one Charlson comorbidity were found to have less than 20% of the odds of death compared with those with ≥3 comorbidities (odds ratio 0.187, $P < .0001$). Patients with two Charlson comorbidities had 0.362 the odds of death compared with the reference group ($P < .0001$). This confirms the need for careful patient selection for CEA.

The strengths of our study lie in its unique approach compared with previous analyses of the volume-outcome effect. We examined the data by inspection of logit-transformed lowess smoothing functions for appropriate cutoffs rather than using quantiles or empiric volume categories. By using the data itself to choose volume categories, we avoided "fishing" for cutoffs rendering the lowest P values, instead choosing the best combination of knots based on log-likelihoods.

The weaknesses of our study mirror those with the use of any secondary analysis of administrative data. We believe the data source to be valid, based on previous corroboration with our institution's CEA data.[29] Nonetheless, as with any administrative dataset, there is the potential for non-differential or differential underreporting of events. Non-differential underreporting would bias our findings towards the null, whereas systematic, differential underreporting could change the direction or magnitude of the effect. Another threat to internal validity of administrative data includes confounding.[35] Our data lacks information on surgeon or hospital characteristics, which could act as unmeasured confounders in the relationship between CEA annual volume and in-hospital death. For instance, our data contains no information on specialty training, years in practice, or other factors that could contribute to technical expertise. The high number of surgeons performing a single CEA annually (214 unique practitioners over the 10-year study, or 48.4%) begs the question of whether these "surgeons" in fact represent referral physicians or trainees. We feel this is unlikely, given our requirement of the presence of Diagnosis Related Group 5 in our case identification algorithm. Given the ghost coding of physicians in the Maryland HSCRC database, it is not possible to know the board certification or fellowship training status of the providers and it therefore remains possible that some non-surgeons or trainees were captured in our analysis.

Additionally, although we used the HSCRC database for 10 years and identified 22,772 patients who underwent CEA, the database included only 123 deaths. This low number of events may have limited our ability to detect statistically significant differences. Our analysis is limited by the inability to differentiate deaths attributable directly to CEA from those related to underlying medical conditions.

Some patients may have died from unrelated causes, such as pulmonary embolism or arrhythmia. The use of a more common complication such as post-CEA stroke would have provided a larger number of events and a more relevant volume-outcome relationship.

We chose to control for patient comorbidity using those conditions identified by Charlson because this method remains the most widely used, and has been validated in multiple languages and for other diseases. However, Elixhauser has developed a method with some strengths over the Charlson approach, by capturing acute illnesses, adding a number of comorbidities, and dropping others that seemed conceptually inappropriate.[36] We also did not use the weights included in the original Charlson Index, as we did not want to assume that these weights derived from Charlson's set of medical patients would apply to the surgical patients in our dataset. Of perhaps more relevance, the Maryland HSCRC database lacks present-on-admission coding. It is therefore possible that some comorbidities may in fact represent complications of the CEA itself. Despite these limitations, we feel that our approach remains a reasonable method to provide some measure of control over comorbidities in this study demonstrating a rigorous statistical treatment of the volume-outcome effect. We elected to control for comorbidity rather than symptomatic status as methods for comorbidity adjustment are well-validated, whereas those for identification of symptomatic status are not. Previous work from our group has demonstrated that the overwhelming majority of patients undergoing CEA today in Maryland, as elsewhere, are asymptomatic. This is consistent with the preponderance of asymptomatic patients treated in most of the very recent carotid stent trials and registries.

We have demonstrated a technique for rigorous statistical analysis of volume-outcome data and have found a modest volume effect for death after CEA in this 10-year dataset from Maryland after controlling for patient comorbidities and age. Higher volume surgeons had lower estimated odds of death, particularly when considering those performing four to 15 CEAs per year. These data suggest that a patient undergoing CEA by a surgeon performing an average of 16 CEAs annually has a statistically equivalent risk of death compared with one undergoing CEA by a surgeon performing any number higher than this, when controlling for hospital volume, patient comorbidity, and patient age. Hospital volume was not seen to be as significant a predictor of post-operative death in this study, with only high volume hospitals (>130 CEAs per year) showing a statistically significant decrease in the estimated odds ratio of death. Patient comorbidity was a significant predictor of in-hospital post-operative death even after controlling for surgeon and hospital volume and patient age, confirming the need for careful patient selection for CEA. As studies on volume-outcome relationships can have important implications for health policy and surgical training, such studies should consider non-linear effects in their modeling of procedural volume.

AUTHOR CONTRIBUTION STATEMENT

Conception and design: SN, GY, RT, MG, DC, BP
Analysis and interpretation: SN, GY, RT, MG, DC, BP
Data collection: SN, GY, DC
Writing the article: SN, GY, RT, DC, BP
Critical revision of the article: SN, GY, MG, BP
Final approval of the article: SN, GY, RT, MG, DC, BP
Statistical analysis: SN, GY, RT, MG, DC
Obtained funding: SN, RT, MG
Overall responsibility: SN

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INVITED COMMENTARY

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Dr. Nazarian and her colleagues have provided an elegant statistical analysis of the carotid endarterectomy (CEA) volume/outcome effect on mortality using a hospital administrative database. Unfortunately, administrative databases are derived from hospital claims which are focused on payment and are of only limited utility for outcome studies. The capture of procedure indication, comorbid conditions, and complications may be incomplete or inaccurate. The inability to accurately stratify CEA patients by indication or validly capture postoperative stroke is an important limitation.

The striking finding that approximately half the surgeons (214/442) in Maryland are performing one or fewer CEA procedures a year is potentially more suggestive of a problem with the database than necessarily an accurate representation of practice. The database is derived from hospital claims not those of the surgeon. The data field used to determine the physician of record for the hospital stay could have non-surgeons (eg primary care physicians or hospitalists) in that field. Acceptance of this finding requires additional validation.

It is hard to argue against the concept that experience with a procedure is generally associated with better results. The question remains, however, whether or not annual volume is a valid surrogate for experience and even more importantly judgment. Is annual volume as good a predictor for the surgeon who has had adequate vascular training and practices 100% vascular surgery, but with a relatively low volume of CEA procedures, as for the surgeon without vascular training whose only vascular procedure is an occasional CEA? These are important questions that are not answered in this study.

I would argue that the perverse incentives associated with using an annual procedure volume threshold alone to direct payment policy are likely to result in more harm than benefit. Overemphasis on procedure volume creates an incentive to do procedures rather than counsel patients adequately about the risks versus benefits of that procedure. In any study that has large numbers, it is important to recognize that statistical significance does not indicate the magnitude of the effect. In the current study, the mortality of CEA for the group of surgeons who did four to 14 procedures a year was 0.7% versus 0.4% for the surgeons who did 15 or more. The benefit of carotid intervention in asymptomatic patients is not high. Any policy that creates more incentive for the potential overuse of carotid intervention in asymptomatic patients is likely to cause more population harm than any benefit accrued from this small absolute decrease in mortality.